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Maladaptive Daydreaming: The German and the Dutch Versions of the 16-Item Maladaptive Daydreaming Scale

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The 16-item Maladaptive Daydreaming Scale is a self-report measure translated into 37 languages, designed to capture the experience of compulsive and excessive daydreaming. In an online study comprised of monolingual and bilingual participants ($N = 201$), we examined the psychometric properties of the German and Dutch versions of the scale. We also explored associations of maladaptive daydreaming with fantasy proneness (i.e., characteristics associated with intensive immersion in fantasy) and counterfactual thinking (i.e., the ability to conceptualize alternative scenarios to reality). We found no differences between both language versions, suggesting equivalence. Also, reliabilities were adequate, and the four-factor structure was replicated. We confirmed previous findings of a positive correlation between fantasy proneness and maladaptive daydreaming ($\rho = .58$, $p < .01$) but found no evidence that maladaptive daydreaming is associated with an increased capacity to generate counterfactuals, although our measure of this capacity was far from optimal. Other study limitations and further research approaches are discussed.

Keywords: maladaptive daydreaming, counterfactual thinking, fantasy proneness, 16-item Maladaptive Daydreaming Scale, psychometrics

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Maladaptive daydreaming (MD) is a proposed psychological disorder that was first defined as an "extensive fantasy activity that either replaces human interaction, interferes with academic, interpersonal, or vocational functioning, or both" (Somer, 2002, p. 197; see also Soffer-Dudek & Somer, 2022; Soffer-Dudek & Theodor-Katz, 2022). Unlike mind wandering and normal daydreaming, there are good reasons to assume that MD is unique in its intensity, underlying mechanism, and consequences. Soffer-Dudek and Oh (2024, p. 1) argued that the core of MD is the tendency to "addictively engage in fanciful, narrative, and emotional daydreaming for hours on end." This tendency is accompanied by reduced self-awareness and impaired attention to the immediate external environment (Soffer-Dudek, 2019).

Normal daydreaming serves various adaptive purposes, including future planning, creativity, problem solving, attentional cycling, and dishabituation and does not necessarily presume pathological levels of fantasy proneness and/or dissociation (e.g., Klinger et al., 2009). Conversely, MD is closely linked to one facet of fantasy proneness, called immersive daydreaming (West & Somer, 2020), a highly rewarding off-task mentation that may initially help with mood enhancement, wish fulfillment, companionship, intimacy, soothing as well as disengagement from boredom, stress, and pain (Somer, 2002). MD typically encompasses intricate imagined scenarios centered around emotionally gratifying themes, like social

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The data that support the findings of this study can be found on https://dataverse.nl, the open-data platform of Dutch Universities: https://doi.org/10.34894/AY87EJ.

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validation and support. These experiences are frequently triggered and maintained by repetitive movements and exposure to music, resulting in a compulsive involvement in vivid fantasies (Soffer-Dudek & Somer, 2022; Somer et al., 2016). Individuals with MD often find it difficult to manage this compulsive engagement, which can significantly disrupt their daily lives (Musetti et al., 2023). Eventually, it may develop into a harmful disorder that resembles behavioral addiction (Pietkiewicz et al., 2018; see also Somer et al., 2020). The general defining factor in favor of categorizing this type of excessive daydreaming as a disorder, separating it from normal daydreaming, includes interference with one's day-to-day life and disruption of long-term goals (Soffer-Dudek et al., 2021). It is not primarily the content of the fantasies but rather the persistent and compulsive immersion in them that leads to the distress and impairment experienced by people with MD. Unsurprisingly, MD is often associated with comorbid conditions like anxiety disorders and depression (Somer et al., 2017). Vyas et al. (2024) astutely pointed out that symptoms of MD bear similarities to rumination and various psychological disorders, particularly attention deficit hyperactivity disorder and dissociative disorders. These and other authors (e.g., Roy et al., 2024) propose the existence of a shared neurophysiological dysfunction that manifests divergently across these conditions.

Regardless of MD's precise connections with attention deficit hyperactivity disorder, dissociative symptoms, and addiction characteristics, fantasy proneness is the most well-established correlate of MD (Bigelsen et al., 2016). In 1983, Wilson and Barber's groundbreaking study identified a group of individuals they termed "fantasy prone persons," characterized by deep and longstanding involvement in fantasy. These individuals reported immersive, vivid, and rewarding fantasy experiences that were intricately woven into their daily lives. Wilson and Barber (1983) argued that fantasy proneness is a stable, trait-like characteristic defined by frequent daydreaming, intense fantasies, and physical reactions to these fantasies (e.g., feeling ill at the thought of rotten food). Subsequent research has largely validated Wilson and Barber's (1983) typology, but unlike MD, fantasy proneness is not inherently maladaptive (Merckelbach et al., 2022). Nevertheless, Merckelbach et al.'s (2022) meta-analysis reviewed five separate studies examining the relationship between fantasy proneness and MD, finding an average correlation of .51, 95% confidence interval [.46, .56], based on an aggregated *N* of 1,593.

One prominent aspect of fantasy proneness is a preference for exploring the limits of reality and convention, which, in principle, is a benign feature closely related to counterfactual thinking (CFT). CFT refers to the ability to conceptualize alternative scenarios to reality (McEleney & Byrne, 2006). Counterfactual thoughts are mental representations of alternatives to events, actions, or states (Byrne, 2005). Counterfactuals are not restricted to past events but may also be directed to the future, involving the production of what has been termed "prefactuals" (i.e., mental simulations of what might happen; Epstude & Roese, 2008). By considering how they could control the outcome, prefactuals help prepare for future behavior via the formation of intentions and in supporting future decisions (Byrne, 2016). Some authors have argued that CFT about past events and prefactual thinking about future events serve different functions. However that may be, people differ in the degree to which they are able to think counterfactually and prefactually (e.g., Bacon et al., 2020).

Bacon et al. (2013) found in their student samples (*N*s = 106 and 76) that fantasy proneness is positively related to the ability to generate counterfactuals (*r*s = .55 and .58, respectively). This finding highlights the potential benefits of fantasy proneness, as it may enhance CFT and thereby creative problem solving. In contrast, West and Somer (2020) utilized the Biographical Inventory of Creative Behaviors to assess creative outputs, such as writing short stories, among individuals with varying degrees of excessive daydreaming. Their findings indicated a decrease in creative output among those exhibiting MD. Thus, while fantasy proneness is associated with both a greater tendency for CFT and excessive daydreaming, it appears that the most extreme forms of daydreaming are linked to diminished creative thinking.

The present study, first and foremost, focused on filling gaps in recent research developments concerning MD and its measurement. The previously proposed German and Dutch versions of the 16-item Maladaptive Daydreaming Scale (MDS-16) by the International Consortium for Maladaptive Daydreaming Research lacked clarity and well-formulated Likert scale anchors. Moreover, the German and Dutch translations were not tested on native-speaking populations, and the psychometrics of these versions are yet to be investigated.

Therefore, we revised the German and Dutch versions of the MDS-16 to improve their clarity and then evaluated the psychometric properties of these versions in samples of native speakers. Thus, we examined the reliability (i.e., internal consistency) of these as well as their convergent validity by assessing correlations with a measure of fantasy proneness.

Drawing on the established connection between fantasy proneness and MD (Somer et al., 2016) and the relationship between fantasy proneness and CFT (Bacon et al., 2013), another aim of the present study was to explore whether MD is related to CFT.

We tested the following hypotheses: (a) The German and Dutch versions of the MDS-16 demonstrate high internal reliability and equivalence; (b) Individuals with high MDS-16 scores will exhibit greater fantasy proneness, as measured by the Creative Experience Questionnaire (CEQ; Merckelbach et al., 2001). Such a finding would replicate earlier results (Bigelsen et al., 2016) and support the convergent validity of the MDS-16 in Dutch and German samples; (c) Individuals who self-identify as having MD will have significantly higher MDS-16 scores than control participants.

Additionally, we explored whether MD scores correlate with an increased or a reduced level of counterfactual thoughts, using the task also employed by Bacon et al. (2013) in their study on fantasy proneness and CFT.

Method

Participants

The standing ethical committee of the Faculty of Psychology and Neuroscience of Maastricht University approved the study (ERCPN 244_146_11_2021). Participants were recruited via different social media platforms, snowballing in close circles, and via the SONA participation system of the Faculty of Psychology and Neuroscience of Maastricht University. Social media platforms included Instagram, Discord,[1] Tumblr (https://madd-information.tumblr.com/), and a maladaptive daydream forum (https://maladaptives-tagtraeumen.forumprofi.de/). Inclusion criteria were consenting age and adequate mastery of German, Dutch, or both languages.

All participants provided informed consent. A total of 27 participants had missing records for our measures of fantasy proneness and/or the CFT vignette (see below); three bilingual participants had missing records for the second half of the MDS-16. After excluding participants with missing records, the sample consisted of 204 participants (76.5% women, 15.7% men, 6.4% nonbinary, 1.5% preferred not to say; mean age range 18–24; 63.2% native German speakers and 38.8% native Dutch speakers; 81.2% with a degree equivalent to or higher than the highest secondary school degree, e.g., Abitur in Germany), including German (n = 85), Dutch (n = 40), and bilingual (n = 79) subsamples. Three participants failed an embedded check for inattentive/random responding while also not completing the CFT task. They were excluded, leaving 201 participants in the final sample. The bilingual subsample was mainly recruited via the SONA participation system and mostly comprised Maastricht University students (SONA participants n = 50). We also advertised the study on German and Dutch Internet communities for individuals with MD. Language proficiency was determined by directly asking participants about their native language(s). It is worth noting that our university is located in a border region where many people, due to factors such as a mixed family background or the city where they grew up, speak both German and Dutch fluently. We asked participants whether they would define themselves as excessive daydreamers and 132 (66%) indicated that they consider themselves as such. The term "excessive daydreamers" was not further described or elaborated on for the participants. However, the proportion of self-reported excessive daydreaming varied across the subsamples: it was relatively low in the bilingual student subsamples (45%) recruited via the online SONA participation system but relatively high (70%) in samples recruited via social media channels. This difference could be due to our promotion of the study as research on excessive daydreaming: the call for participants may have attracted people who suffer from this condition but other variables (e.g., education status, demand characteristics) may also have played a role.

[1] Discord Server: Daydreamer's Hideaway.

Materials

Demographics, specifically native language, age, gender, education status, and career, were surveyed, followed by daydream context and content questions. Additionally, we asked participants if they consider themselves to be excessive daydreamers. The Dutch and German versions of the MDS-16 (Somer et al., 2016), the CEQ (Merckelbach et al., 2001, 2022), and a vignette prompting counterfactual thoughts drawn from McEleney and Byrne (2006) were employed. All were administered in German or Dutch via Qualtrics.

The 16-Item Maladaptive Daydreaming Scale

The MDS-16 developed by Somer et al. (2016) is a self-report instrument designed to capture the experience of compulsive immersive daydreaming. It includes 16 items that cover four underlying dimensions (Soffer-Dudek et al., 2021; Uslu, 2015): (a) impairment, which highlights the disruption and dysfunction individuals experience associated with daydreaming; (b) music, which is employed to trigger and enhance the immersive fantasy; (c) kinesthesia, which reflects repetitive body movements to initiate or prolong daydreaming; and (d) yearning, which refers to the addictive nature of daydreaming. Respondents are asked to answer the items on a 0% (i.e., never) to 100% (i.e., always) scale. Scores are averaged to obtain a mean MDS-16 score. Ross et al. (2020) summarize evidence to show that the English version possesses good criterion-related validity and adequate test–retest reliability. Although there is some discussion in the literature concerning the optimal cutoff for the scale (e.g., Abu-Rayya et al., 2019), in the present study, we followed Soffer-Dudek's (2021) recommendation and used a cutoff of 40, a mean score reported to best discriminate between controls and cases of self-diagnosed MD, with a sensitivity and specificity > 90%.

We developed new Dutch and German versions of the original MDS-16 utilizing back translations. This technique involves multiple translators who translate from the original language to the target language and then back from the target language to the original, ensuring that the translated version retains the same meaning as the original. Additionally, two experienced clinicians fluent in German or Dutch provided detailed comments on the German and Dutch draft translations. This procedure resulted in further adaptations. Notably, much attention was paid to the anchors of the scale. The initial translation included more ambiguous terms (e.g., "very often," "Sehr häufig"). We followed the suggestions by Schwarz (1999) by choosing more strongly worded anchors that match the numerical extremes "0%" and "100%." Consequently, in the final version, we replaced the MDS-16 anchors with more radical qualifiers such as "altijd" in Dutch or "immer" in German, both meaning "always." The Dutch and German versions of the MDS-16 can be found in the Supplemental Materials.

The Creative Experience Questionnaire

The CEQ (Merckelbach et al., 2001) is a 25-item scale with a yes–no format. The scale aims to capture the essence of fantasy proneness by examining key aspects, including developmental antecedents, intense engagement in fantasy, and the outcomes of fantasizing. Affirmative responses are summed to obtain a total score, with CEQ scores above 12 indicative of high fantasy levels (cf. Merckelbach et al., 2022). The CEQ was chosen due to its brevity compared to other measures (e.g., the Inventory of Childhood Memories and Imaginings; Wilson & Barber, 1983) and its satisfactory psychometric properties. The questionnaire shows adequate test–retest stability and internal consistency (Merckelbach et al., 2001). The CEQ was used as the main criterion for the convergent validity of the MDS-16 in its original validation study (Somer et al., 2016). That study found the MDS-16 to be significantly, but not perfectly, associated with the CEQ ($r = .58$, $p < .01$), indicating that both constructs are related yet distinct.

Counterfactual Thinking (Vignette)

A case vignette was employed to measure individual differences in CFT. The original version by McEleney and Byrne (2006) was translated into German and Dutch. The vignette is about a person who moves to another city for a new job and feels lonely in the first weeks after the move. Participants read the vignette and were invited to identify with the main character. The instruction was as follows: "Please read the scenario and imagine that it really happened to you. Then write about the imagined experience as if you were writing in your diary. Include your thoughts and feelings about all the events and the outcome" (see also, Bacon et al., 2013, 2020). Next, they were given 5 min to write a fictional diary entry based on the vignette supplied. Two independent judges (Pia Breuer and Harald Merkelbach) evaluated entries to determine the number of counterfactual thoughts using the guidelines provided by Bacon et al. (2013). They defined CFT as "any thought about how a change to the scenario would change the outcome" (p. 470). An example of a counterfactual thought would be: "I should have visited my aunt who lives in this city; she knows many people, and through her, I could have gotten to know them too." Bacon et al. (2020) found a high interrater agreement in evaluating the presence of counterfactuals. In our study, the interrater agreement between the two judges was poor ($\kappa = .32$) and so frequencies of the two raters were averaged. Additionally, the word count of the diary entry was considered.

Procedure

This online study was conducted using the Qualtrics survey platform and involved three groups differing in language. Following the approach by Giger and Merten (2019) and Dandachi-FitzGerald et al. (2023), the first group, bilingual participants, were given a version of the MDS-16 that contained half German and half Dutch items. Based on random allocation, this bilingual group either received the first eight items in German and the second eight items in Dutch (the Deutsch Nederlands subsample) or vice versa (the Nederland Deutsh [NLDE] subsample). This way, the equivalence of Dutch and German versions of the MD scale could be tested. The MDS-16 was presented fully in German or in Dutch in the two non-bilingual groups. Participants in the bilingual condition were asked about their native language and to confirm their proficiency in both languages before continuing the study. Within the two non-bilingual conditions, participants skipped this step and directly moved on to a brief set of questions about their demographic background. This set of questions included a singular item checking for self-identification as an "excessive daydreamer" (e.g., item in German: "Halten Sie sich für eine*n übermäßigen Tagträumer*in?"). Following the MDS-16, the CEQ and vignette were displayed in the participants' native or more proficient language. All conditions included an attention check. This check involved an item that asked participants to select 80% as a response on a visual analogue scale (e.g., attention check instruction in German: "Um zu zeigen dass, Sie aufmerksam sind, wählen Sie bitte '80%' auf dem Schieberegler als Antwort aus.") and three participants failed it while also not producing a diary entry during the CFT task; consequently, their records were removed from the study, leaving 201 participants in the sample. To maximize statistical power, participants who failed the attention check but still produced an output during the CFT task were still included in the analysis. Participants completed the survey in about 20 min, and we rewarded students at the Faculty of Psychology and Neuroscience at Maastricht University with 0.5 SONA credits. Other participants were not compensated.

Statistical Analysis

To ensure high-quality psychometric data for the German and Dutch versions of the MDS-16, we adopted a stringent approach to handling noisy data: any individual with questionable responses to the attention-check item combined with a lack of output on the vignette task was excluded from the final data set. For participants with missing data, we did not follow a listwise deletion procedure, but considered each measure, separately. For example, we had 201 complete records for the MDS-16, but only 177 complete records for the CFT task. This explains the variable degrees of freedom below. For the preliminary data analyses, we used IBM SPSS Statistics 27 (IBM Corp., Armonk, New York). First, we obtained descriptive statistics of the various subsamples and calculated Cronbach's alphas for the MDS-16. Second, following the procedure of Giger and Merten (2019), we tested the equivalence of the scale in both languages, employing independent and paired sample *t*-tests to compare the two language versions. Third, utilizing Mplus (Version 7.2), we conducted a confirmatory factor analysis (CFA) based on the four factors proposed by Soffer-Dudek et al. (2021). Given the language equivalence established in the prior analysis, we collapsed language versions, combining data from all groups to ensure a sufficiently large sample size for a CFA, resulting in a combined MD scale data set of $N = 201$. The examined model revealed the following structure: Factor 1 (F1) represented the Impairment component, comprising Items 5, 6, 7, 8, 9, and 11. Factor 2 (F2) depicted Music and included Items 1 and 16. Kinesthesia was captured by Factor 3 (F3), with Items 3 and 14 loading onto it. Finally, Factor 4 (F4) encapsulated Yearning, encompassing Items 2, 4, 10, 12, 13, and 15. Subsequently, factor and sum scores for each of the four subscales were calculated, and corresponding reliability analyses were performed.

Additionally, as the interrater agreement for CFT scores turned out to be low, we averaged the CFT scores of the two raters. Next, we calculated Spearman's rank correlation coefficients between MD, fantasy proneness, and CFT ability (measured with the MDS-16, CEQ, and the CFT vignette, respectively). Finally, using *t*-tests and accuracy parameters (i.e., sensitivity and specificity), we compared the MDS-16 scores of those who self-identified as excessive daydreamers and those who did not.

Results

Datafile and background information can be found on https://dataverse.nl, the open-data platform of Dutch Universities: https://doi.org/10.34894/AY87EJ. Table 1 summarizes the MDS-16 data. The average MDS-16 score for the total number of complete cases (N = 201) was 43.7 (SD = 21.9), which is in line with what Schimmenti et al. (2020) reported for their Italian sample that also included a subsample of excessive daydreamers (N = 468; M = 46.68, SD = 22.35). Using the cutoff of 40, we identified 107 participants (53.2%) as having probable MD.

A one-way analysis of variance indicated that the four groups significantly differed with regard to MDS-16 scores, $F(3, 197) = 15.48$, $p < .001$, which perhaps reflects the fact that the DENL and NLDE subsamples consisted of university students. In contrast, the other groups had a more heterogeneous demographic composition and were recruited through social media platforms, including one dedicated to MD. Post hoc Tukey tests revealed that the German and Dutch subgroups did not differ, whereas those between the two bilingual subgroups were marginal (see below). However, significant differences were evident when comparing Dutch and German subgroups against the bilingual groups (all ps < .01).

We inspected demographic variables (age, gender, and educational background) and MDS-16, CEQ, CFT and self-defined excessive daydreamers scores across the full sample. Results are summarized in Supplemental Table 1. Only associations with gender reached significance. Bonferroni corrected follow-up comparisons revealed a significant difference between men (n = 32) and nonbinary individuals (n = 13) for MDS-16 scores. Nonbinary individuals scored significantly higher compared to male participants. Even though these results may display an impact of gender on the MDS-16 scores, important of note here is the small sample size of both groups, which may have biased the results. For the CEQ scores, significant differences between all three gender categories could be found. Nonbinary individuals displayed the highest mean scores, followed by women and then men. This study is the first to report a gender difference on the CEQ.

Supplemental Table 2 provides an overview of associations between context and content of daydreaming and scores on MDS-16, CEQ, CFT, and self-defined excessive daydreamers. Correlations that reached significance were of modest size, suggesting that MD tendencies are associated with work and social interaction contexts, as well as with movie and TV series content.

Table 1
16-Item Maladaptive Daydreaming Scale Scores of Total Sample and Subsamples

Group	*N*	*M*	Range	*SD*	Cronbach's α
DE	85	50.6	6.3–97.5	23.9	.94
NL	40	50.5	8.3–81.3	20.3	.93
DENL	38	28.1	7.9–54.4	13.7	.83
NLDE	38	34.5	5.9–57.3	13.2	.83
Total	201	43.7	5.9–97.5	21.9	.89

Note. Internal reliability coefficients (Cronbach's α) are also shown. DE = German; NL = Dutch; DENL = Deutsch Nederlands, first half German, second half Dutch; NLDE = Nederland Deutsh, first half Dutch, second half German.

Reliability and Equivalence

Table 1 provides Cronbach's alphas (αs) for the total and subgroups. All αs were >.80, indicating excellent internal reliability. Three observations regarding the equivalence of German and Dutch MDS-16 versions are of note. First, the subgroups that only completed the German or Dutch version attained similar means and standard deviations, $t(123) < 1.0$, ns. Second, the "German half first" subgroup (DENL) attained lower scores than the "Dutch half first" subgroup (NLDE), $t(74) = 2.07$, $p = .04$, two-tailed, with a medium-small effect size (Cohen's $d = 0.48$). Third and most importantly, we compared the equivalence of means within each bilingual subgroup with paired *t*-tests. For the DENL subgroup, means and *SD*'s for first and second half were 29.9 ($SD = 17.2$) and 26.3 ($SD = 13.7$), respectively, $t(36) = 1.49$, $p = .15$. For the NLDE subgroup, these scores were 34.0 ($SD = 16.1$) and 35.0 ($SD = 13.9$), respectively, $t(36) = 0.42$, $p = .68$. Overall, this pattern suggests that scores were not dependent on language version.

Confirmatory Factor Analysis

Table 2 summarizes the results of the CFA. By default, Mplus makes use of the full information maximum likelihood estimation method, enabling the analysis of all available data. All standardized factor loadings are >.45. The lowest loading observed (.47) is the association of Item 15 with F4. Overall, Table 2 provides empirical support for the four-factor structure proposed by Soffer-Dudek et al. (2021). Table 3 describes the goodness-of-fit measures for the model tested. With comparative fit index and Tucker–Lewis index both around .95, root-mean-square error of approximation (.07) < .08 and standardized root-mean-square residual (.05) < .08, all parameters

Table 2
Confirmatory Factor Analysis

Factor/Indicator	Standardized loading	Error variance
Factor 1: Impairment		
MD5	.83	.32
MD6	.82	.33
MD7	.73	.47
MD8	.87	.25
MD9	.65	.58
MD11	.87	.24
Factor 2: Music		
MD1	.92	.15
MD16	.56	.68
Factor 3: Kinesthesia		
MD3	.74	.46
MD14	.82	.33
Factor 4: Yearning		
MD2	.67	.55
MD4	.80	.36
MD10	.73	.47
MD12	.79	.38
MD13	.64	.59
MD15	.47	.78

Note. Confirmatory factor analysis on 16-item Maladaptive Daydreaming Scale scores of the full sample ($N = 201$). MD = maladaptive daydreaming item.

Table 3
Goodness-of-Fit Measures for Four-Factor Solution of 16-Item Maladaptive Daydreaming Scale

$\chi^2(df = 98)$	CFI	TLI	RMSEA	SRMR
186.14 $p < .001$	.95	.94	.07	.05

Note. All fit measures were based on maximum likelihood estimates. A good fit is indicated by CFI/TLI > .95, RMSEA < .08 (with <.05 indicating close fit) and SRMR < .08. CFI = comparative fit index; TLI = Tucker–Lewis index; RMSEA = root-mean-square error of approximation; SRMR = standardized root-mean-square residual.

indicate a good fit. Several items indicated normality violations in the form of negative kurtosis values between −1.50 and −1. Using a robust estimator (maximum likelihood with robust standard errors; MLR) corrects for these normality violations. These corrected results translated into a further improvement of the fit measures, with chi-square ($df = 98$) leveling off to 165.6, comparative fit index and Tucker–Lewis index remaining constant at 0.95 and 0.94, respectively, and root-mean-square error of approximation declining to .06. standardized root-mean-square residual was not affected by the robust estimator.

Table 4 shows the correlational pattern of the four different factors. The highest correlation (.84) was that between F1 and F4, encompassing Impairment and Yearning, respectively. The lowest correlation (.49) is that between F1 and F2 (Impairment and Music). Figure 1 graphically summarizes the factor loadings of each item and the correlations between the factors in the four-factor model.

A Likelihood Ratio test comparing our model (where all factor loadings are freely estimated) with a model that imposes an equality constraint on all loadings within factors yields a chi-square value of 73.88 with 12 degrees of freedom. This result is highly significant ($p < .001$), showing that tau equivalence of items within factors cannot be defended. However, in the case of larger samples chi-square tests are known to be sensitive and may flag unimportant discrepancies as significant. To further investigate the case for equal factor loadings, factor scores for each of the four factors and sum scores for each subscale were computed. Supplemental Table 3 shows that the factor scores highly correlate (all > 0.9) with the corresponding sum scores, providing an argument (contrary to

Table 4
Correlation Between Four Factors of the 16-Item Maladaptive Daydreaming Scale

Factor	1	2	3	4
1. F1	—	.49	.71	.84
2. F2		—	.61	.59
3. F3			—	.82
4. F4				—

Note. F1 = Impairment; F2 = Music; F3 = Kinesthesia; F4 = Yearning.

the significant chi-square difference test) that simple sum scores can be used for assessing individuals in terms of MD dimensions. As shown in supplemental Table 4, the sum scores for each of the subscales correlate highly with the total sum scores (based on all 16 items taken together). This provides an argument for considering the total score as an informative measurement of MD (apart from the more detailed information given by the four separate sum scores). Supplemental Table 5 summarizes the reliability analysis of the sum and total sum scores. All reliabilities range from acceptable (for the sum score of the Music dimension—comprised of just two items, which makes it less reliable) to excellent.

Counterfactual Thinking

Table 5 summarizes the CFT data. As a poor interrater agreement (κ = .32) between the two judges was observed, a combined score was calculated and employed for further analysis. The average length of participants' diary entries was 101.85 words. They contained a mean of 0.75 counterfactual thoughts per participant, ranging from 0 to 6.5 per person (SD = 1.02). The mean number of counterfactuals (M = 0.75) was comparable to the mean frequency (M = 0.79; SD = 1.05) obtained by McEleney and Byrne (2006; Experiment 1), but lower than that found by Bacon et al. (2013; Experiment 1; M = 117; SD = 1.32).

Correlations With CEQ and CFT

Table 6 shows Spearman's rank correlation coefficients across the entire sample between MDS-16, CEQ, and parameters of the CFT task. MDS-16 scores were significantly and positively related to CEQ. MDS-16 scores and number of counterfactuals were not related to each other. Unlike previous studies (Bacon et al., 2013), there was no significant association between CEQ and number of counterfactuals. The word count parameter was significantly correlated only with the number of counterfactuals.

Excessive Daydreamers

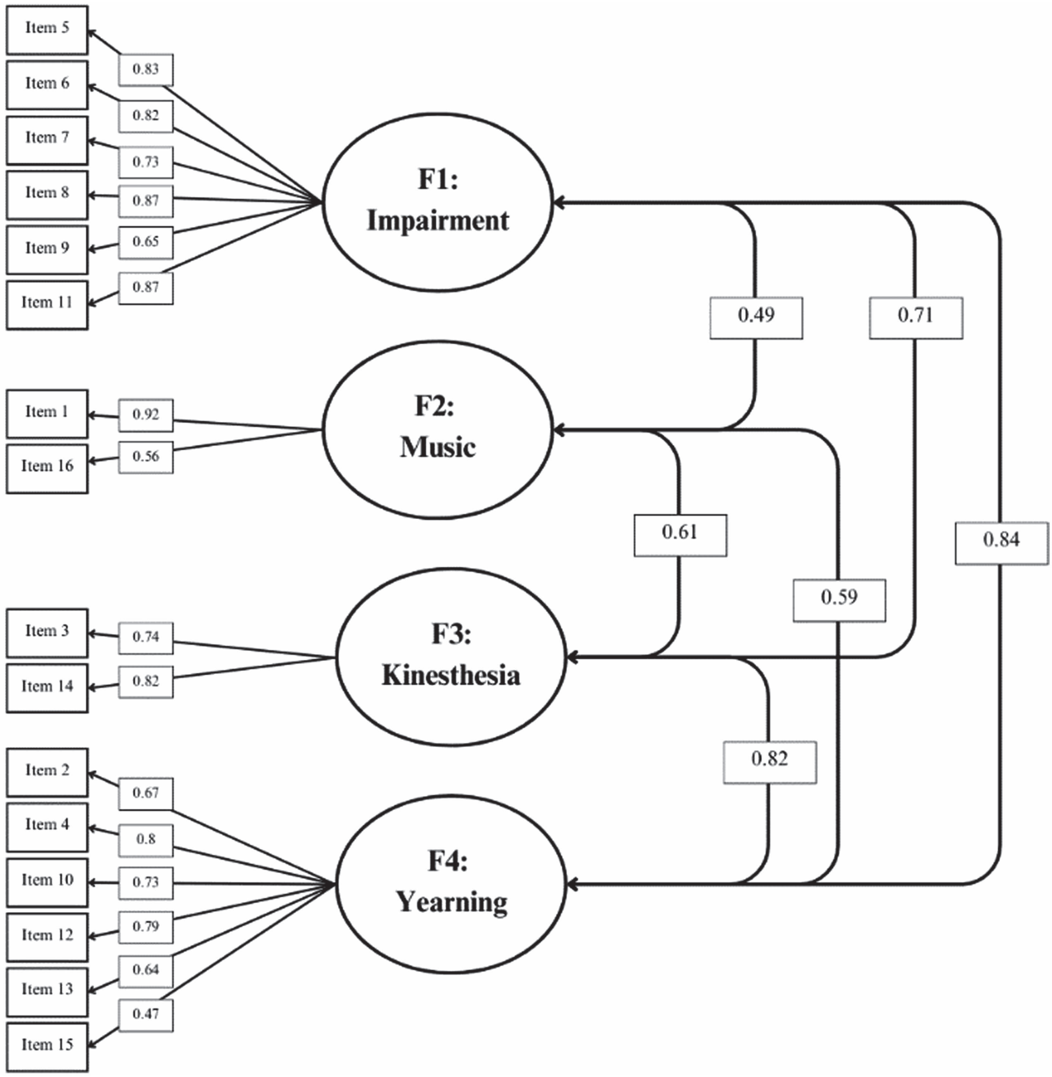
Of the participants who self-identified as "excessive daydreamers" (n = 132), 71.2% had MDS-16 scores above the cutoff of 40 (i.e., sensitivity = 71.2%). Of those who did not self-identify as "excessive daydreamers" (n = 69), 13 (18.8%) attained a score above the cutoff and, therefore, were "misclassified" as maladaptive daydreamers ("false positive rate" = 18.8%). The group that self-identified as "excessive daydreamers" displayed higher MDS-16 scores than those who did not identify with that label, M = 52.6 (SD = 19.7) and 25.4 (SD = 13.1), respectively, $t(199) = 10.33$, $p < .01$, Cohen's d = 1.63. The people who self-reported excessive daydreaming also presented higher CEQ levels, suggesting elevated fantasy proneness in this group: 12.9 (SD = 4.02) versus 8.4 (SD = 3.3), $t(196) = 7.83$, $p < .01$, Cohen's d = 1.22.

Discussion

We investigated the psychometric properties of two language versions of the MDS-16, a measure of MD. In the setting of an initial explorative investigation of the Dutch and German MDS-16, our data display encouraging results. These language versions achieved good internal consistencies, and we observed no language dependence. Furthermore, we confirmed the four-factor model proposed by Soffer-Dudek et al. (2021) for our merged sample of Dutch and German individuals. Our analysis indicates that simple sum scores are suitable for assessing individuals on the MDS-16 dimensions. During recruitment, we specifically targeted individuals from the two linguistic subpopulations who self-identified as excessive daydreamers, and we found that they had considerably higher MDS-16 scores than a comparison group of university students. The majority of excessive daydreamers (71.2%) scored above the MDS-16 cut point.

While translating the MDS-16, we paid careful attention to the scale's anchors and followed the approach outlined by Schwarz (1999), who emphasized the importance of using clear and

Figure 1
16-Item Maladaptive Daydreaming Scale Four-Factor Model



Note. Factor loadings and correlations of the confirmatory factor analysis for the four-factor model.

straightforward labels for anchors. Even though we chose anchors with more significant impact than the ones used in the English MDS-16, no noticeable difference in mean scores emerged when we compared our overall data with those reported for the original versions. However, our changes could have possibly lead to the discrepancy between self-identified excessive daydreamers and detected participants with probable MD. In our sample, 28.8% (n = 38) of participants who labeled themselves as "excessive daydreamers" were missed by using the cutoff of 40 on the MDS-16. Our more strongly worded anchors could have had an impact on people's scores by possibly raising the standards. This change might explain why the scale did not detect a sizeable minority of self-identified respondents with excessive daydreaming in the present study. Alternatively, individuals may be imprecise in their self-definition of excessive daydreaming, a possibility that is also consistent

Table 5
Descriptive Statistics of the Counterfactual Thought Measure

Rater	*N*	Range	*M*	*SD*
PB	177	0–9	.97	1.42
HM	177	0–4	.58	0.82
CCFT	177	0–6.5	.75	1.02
Word count	177	0–1,641	102	134

Note. PB = counterfactual thought score by rater Pia Breuer; HM = counterfactual thought score by rater Harald Merkelbach; CCFT = Combined counterfactual thought score.

with the nontrivial group of people who scored above the MDS-16 cut point but identified as nonproblematic daydreamers (“false positives”). Still another possible explanation is that people are reluctant to disclose that they are coping with excessive daydreaming. This type of underreporting might have played a role in the student sample. The phrasing of the self-identifying item may have contributed to this. During the screening, we used the term “excessive daydreaming” to mitigate the stigma associated with the word “maladaptive.” However, the qualifier “excessive” may be viewed as a pejorative modifier that could have biased respondent self-selection and potentially skewed their responses through demand characteristics or other expectancy effects. Additionally, we did not provide a definition for this qualifier. Therefore, we cannot rule out that this term may have fostered a selection effect.

Table 6
Spearman’s Rank Correlation Correlations Between MDS-16, Fantasy Proneness (CEQ), Counterfactuals (CCFT), and CCFT Word Count

Correlate	MDS-16	CEQ	CCFT
CEQ			
Spearman’s ρ	.58**		
N	198		
CCFT			
Spearman’s ρ	.00	−.04	
N	177	177	
CCFT word count			
Spearman’s ρ	−.05	.03	.38**
N	177	177	176

Note. MDS-16 = 16-item Maladaptive Daydreaming Scale; CEQ = Creative Experience Scale; CCFT = Number of counterfactuals; CCFT word count = Number of words on counterfactual thinking task.
** $p < .01$.

On the other hand, if feelings of shame were an important antecedent of underreporting excessive daydreaming, one would expect a negative correlation between experiences of shame and MDS-16 scores. So far, no such negative correlation has been found. In fact, Ferrante et al. (2022) found a positive and significant correlation of $r = .43$ ($N = 162$).

Given these considerations, the sensitivity and false positive rates we found for the MDS-16 should be interpreted with caution. Future studies on the German and Dutch versions of the MDS-16 should incorporate the clinician-administered Structured Clinical Interview for MD (Somer et al., 2017), which is better suited than self-diagnosis to provide a standard against which the accuracy of the MDS-16 can be evaluated. More generally, follow-up studies are needed to administer the Dutch and German MDS-16 to clinical samples with an already well-established symptom profile of pathological daydreaming. A clinical sample approach might provide a better estimate of the diagnostic accuracy of our MDS-16 versions.

Correlates of Maladaptive Daydreaming

Our data indicate that based on the MDS-16 score, MD is related to fantasy proneness, replicating previous findings (e.g., Bigelsen et al., 2016). That is, we found the MDS-16 to be significantly ($\rho = .58$, $p < .01$), albeit not perfectly, correlated with the CEQ. This further confirms that both constructs are related yet distinct. Furthermore, individuals scoring high on the MDS-16 were not more likely to employ CFT. We have doubts about the informational value of this null finding. Specifically, we were unable to replicate Bacon et al.’s (2013) finding of a positive correlation between fantasy proneness (measured by the CEQ) and counterfactuals (measured with the same task used in this study). Not only did we fail to replicate this correlation, but we also observed low interrater agreement for evaluating counterfactuals, suggesting possible variability in how the CFT task was understood. Despite following Bacon et al.’s (2013, 2020) instructions, the interrater agreement remained poor. Both raters reported difficulties with the provided definition for scoring counterfactuals in the diary entries. Notably, several participants indicated that they either could not complete the task (e.g., due to attention deficit hyperactivity disorder) or did not fully understand the instructions. Thus, it appears that the vignette and scoring instructions offer an unreliable measure of individual differences in CFT. Given these limitations, our data cannot support any conclusions about the relationship between counterfactuals and MD. Future studies should consider developing alternative measures of CFT with improved interrater reliability and re-examining the link between counterfactuals and MD.

Additionally, our results indicate a correlation between gender and scores on the MDS-16 and CEQ respectively. The observed gender differences should be considered with caution in the light of the unequally balanced and small sample sizes of males and nonbinary individuals. Especially considering that for the MDS-16 scores, no significant difference between male and female participants occurred. Studies with larger sample sizes did not report any gender differences. Therefore, our results might simply reflect a biased sample. However, most studies investigating the MDS-16 self-screening tool only report demographic data to describe the sample and neglect to mention if correlational analysis had been performed. Additionally, most studies only report binary genders, neglecting to include other identities. A definite conclusion to gender differences on MDS-16 scores cannot be made. Future studies with larger sample sizes should further investigate if binary versus nonbinary gender differences may occur in MD and fantasy proneness.

Additional Limitations

Additional limitations in the present study need to be acknowledged. First, to conduct a CFA, we combined the four subsamples. As explained above, there is a rationale for this approach given our data; however, the nonsignificant findings that support combining the data are based on relatively small samples and do not strongly demonstrate equivalence between groups. Ideally, we would have conducted an equivalence test to assess this, but such a test requires a much larger sample than we had available. Alternatively, a multigroup analysis could have been used to test the equivalence of factor solutions across subgroups. However, like equivalence tests, multigroup analyses typically require a much larger sample size, with a common guideline of 100 observations per group.

Second, we focused on recruitment among university students for the bilingual scale versions. In contrast, we directed the recruitment for the German and Dutch-only versions primarily toward self-identifying persons with excessive daydreaming, targeting specific forums and online communities. This discrepancy caused by the sampling procedure limits the overarching conclusions drawn from this study. Third, as already mentioned, the self-definition of ordinary versus excessive daydreaming might be imprecise due to subjectivity. Fourth, we used Qualtrics for administering the tests and included only a singular attention check. Therefore, we do not know how attentive participants were throughout the study to the task at hand.

Future Research

Viewing the presented results as an initial, explorative pilot study, we must acknowledge the need for further, more detailed analysis of the psychometrics of the MDS-16 to validate the Dutch and German language versions. We recommend multigroup CFA on a larger sample to expand on our classical test theory approach and further solidify our tentative results. For a thorough examination of the structure and psychometric properties of the MDS-16, it is advisable to employ modern test theory approaches, specifically item-response-theory testing using Rasch models (e.g., Lange, 2017). This will ensure more robust and reliable results. Future studies might also want to investigate if bilingualism impacts the self-classification as “excessive daydreamers” and scoring on the MDS-16 scale. Bilingual samples from languages with distinct grammatical and lexical origins might yield different results. Last, as mentioned above, future studies should consider testing the relationship between CFT and MD with an alternative validated measure of CFT to rule out methodological issues as a cause for null findings.

Conclusion

Our study is the first essential step in demonstrating the psychometric adequacy of the Dutch and German MDS-16. We replicated the previously discovered link between MD and fantasy proneness. Our data provide suggestive evidence that frames the MDS-16 as an appropriate measure to screen for individuals with probable MD.

Though MD is a validated clinical category, it is still underdiagnosed. We believe further research is essential to support these individuals in need. Given that the MDS-16 offers valuable insights for clinicians, we feel that the time is ripe for a large-scale clinical study using the Dutch and German versions of the MDS-16 to further establish its clinical relevance.

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